Parental Gender Preference in the Balkans and Scandinavia: Gender Bias or Differential Costs? *

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Abstract

There is much research indicating the presence of a parental preference for a particular gender of children. The main objective of this paper is to test between the two main explanations for the existence of such preference, namely differences in the costs of raising sons and daughters versus pure gender bias (corresponding to parental utility derived from a child's gender or from children's characteristics exclusive to that gender). First, we use recent EU-SILC data to confirm that in several Balkan and Scandinavian countries the gender of the firstborn predicts the likelihood of a given family having three children or more — a common measure of parental gender preference. Specifically, we confirm son preference in considered Balkan countries and daughter preference in Scandinavian countries. Both having a first child of the preferred gender and of the more costly gender can decrease the probability of having three or more children because parents may already be content or may lack sufficient resources, respectively. Next, we use information on household consumption to differentiate the two explanations. We argue that under the differential cost hypothesis, parents of children of the more costly gender should spend more on children goods and spend less on household public goods as well as on parental personal consumption. In contrast, having children of the preferred gender should increase spending on household public goods since such marriages generate higher surplus and are more stable. Our evidence corroborates the cost difference explanation in countries exhibiting daughter preference.

JEL codes: J13, J16, O15

Keywords: gender preference, gender differences, parental influence, household expenditure

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1 Introduction

The impact of the gender of the first-born child on the number of children in a family has been repeatedly observed in different countries. We confirm son preference using the parity three progression method applied to a pooled EU-SILC 2004-2015 cross-sectional sample from four Balkan countries: Bulgaria, Croatia, Slovenia, and Republic of Serbia [1]. We also confirm the daughter preference for three Scandinavian countries, i.e. Denmark, Norway, and Sweden, which was reported earlier by Andersson et al. (2006); Hank and Kohler (2000). Two possible causes of the gender preference considered in the literature are parental bias in favor of some gender and different expenses of raising sons and daughters (Ben-Porath and Welch, 1976; Lundberg, 2005). In our paper, we aim to find out which of the two is more prevailing in Balkan and Scandinavian countries. Each explanation implies a distinctive relationship between the gender of children and the allocation of household resources. And we test between the two explanations by checking which relationships hold for the household-level data.

Specifically, we find that in Balkan countries households with more female children replace furniture less frequently than households with fewer female children. Moreover in households with more female children mothers report lower ability to spend on oneself. Also, for Balkan countries we find no difference in parental investment in male and female children and no impact of the gender composition of children on ability to make ends meet or the minimum amount of money to make ends meet. We argue based on earlier studies that these findings are consistent with the gender bias explanation and not with the different expenses explanation. At the same time, for Scandinavian countries we find no impact of the gender composition of

[1] these countries are covered by EU-SILC and had the highest SIGI son bias component in Europe according to OECD: https://www.genderindex.org/ranking/sonbias/

children on replacing furniture as well as on consumption of other household public goods. Still, for Scandinavian countries we find significantly larger parental investment in households with more female children. Moreover, we do not find systematic impact of the gender of children on consumption of parents. We argue based on conclusions in Lundberg (2005) and Lundberg and Rose (2003b) that these findings are not consistent with the gender bias explanation but are in line with the differential expenses explanation. Supplementary analysis of the top income decile sub-sample and of cross-country relationships between the gender preference, parental investment, and conventional measures of gender equality supports our point.

2 Literature review

The impact of the gender of the first-born child on the number of children in a family has been repeatedly observed in different countries. The evidence pertains to developing economies (Barcellos et al., 2014; Jiang et al., 2016; Altindag, 2016) and developed economies (Dahl and Moretti, 2008; Andersson et al., 2006; Pollard and Morgan, 2002; Brockmann, 2001). Authors attribute this impact to the parental preference for the gender of children. In developing economies, parents are usually inclined to have more children (progress to higher parities) when their firstborn is a daughter (Filmer et al., 2009; Arnold, 1992). The interpretation of such behavior is that they have the son preference so they continue bearing children until they reach a desired number of sons or the upper limit of the family size. At the same time, in some developed economies parents exhibit son preference (Dahl and Moretti, 2008; Choi and Hwang, 2015) while in other - daughter preference (Andersson et al., 2006; Brockmann, 2001)[2]. The consequences of the parental gender preference have

[2]Sandstrom and Vikstrom (2015) provide evidence for the existence of the son preference in Germany in the second half of the 19-th century which started to fade away thereafter while Outram (2015) finds evidence for the son preference in Edwardian England.

been mostly researched for developing economies. The main consequence is that girls on average have more siblings and receive a lower share of household resources (Vogl, 2013; Jensen, 2003; Basu and De Jong, 2010). Among other consequences are shorter breastfeeding period for girls (Jayachandran and Kuziemko, 2011), worse health and nutritional status of girls (Arnold, 1992), and biased sex ratios (e.g., Jayachandran, 2017; Guilmoto and Duthe, 2013). Some authors also consider consequences of the parental gender preference pertinent to more developed economies. In particular, Kippen et al. (2006) and Dahl and Moretti (2008) argue that the son preference increases fertility in Australia and US respectively. Moreover, Edlund (1999) demonstrates theoretically that the gender preference combined with availability of a gender selection technology [3] could lead to arising of a female "under-class" because poorer parents would prefer having daughters and richer - sons (Trivers and Willard, 1973). Other possible consequences in the setting developed by Edlund (1999) are existence of a "backlog" of unmarried men (Gupta, 2014) with ensuing consequences such as polygamy (Eco, 2018; Seidl, 1995). That is because changes in socio-demographic structure lead to "adoption of adequate institutions" Seidl (1995), which is evident, e.g., in the falling marriage-market value of young men in some localities in US (Autor et al., 2017) accompanied by rising acceptance of polygamy in US (Eco, 2018). Any effective policy that mitigates the effects of gender preferences would need to take into account the causes behind the observed behavior (Lundberg, 2005). Two possible causes considered in the literature are parental bias in favor of some gender and different costs of raising sons and daughters (Ben-Porath and Welch, 1976; Lundberg, 2005). In our paper, we aim to find out which of the two is more prevailing across selected European countries. Each explanation implies a distinctive relationship between the gender of children and the allocation of household resources. And we test between the two explanations

[3] It could, however, still be infanticide, sex-selective abortion, or poorer health care.

by checking which relationships hold for the household-level data.

Regarding the parental gender bias, it has several related definitions in the economic literature. The first definition is that some gender brings more direct utility than the other one or has a utility premium. This definition is used in most papers on the subject (e.g., Jayachandran and Kuziemko, 2011; Dahl and Moretti, 2008; Yoon, 2006). The authors either forgo explaining it and take the gender-biased fertility behavior as their starting point (Jayachandran and Kuziemko, 2011) or explain it by tastes (Dahl and Moretti, 2008) or cultural and biological factors (Yoon, 2006). Scholars in demographic and sociological literature elaborate more on this matter and offer further explanations for the gender bias, such as expansion of the self, affiliation, stimulation, accomplishment or social comparison (Hank, 2007) along with emotional value of children (Sandstrom and Vikstrom, 2015). Moreover, mothers and fathers can perceive the extent to which sons and daughters posses these characteristics differently (Hank, 2007). Finally, the definition proposed in Lundberg (2005) encompasses the aforementioned elements stating that 'parents have child-gender preferences if the marginal value of an additional male child differs, *ceteris paribus*, from the marginal value of an additional female child, or if the marginal utility of increments in boy quality is not equal to the marginal utility of girl quality.' Here 'quality' means child outcomes that are outputs of a household production process where inputs are parental time and market goods and services. This definition incorporates two different cases. In the first case, parental valuation of the gender of children or accompanying outcomes does not relate to parental outlays on children (beyond providing for minimal subsistence level). And in the second case children's outcomes are closely dependent on parental inputs until these inputs reach significant values. The second case is not consistent with previous definitions since the gender is not preferred *per se* but because it makes cheaper the technology of producing some quality, i.e. it is only one of means to reach a specific discrete end.

And in this paper we understand gender bias as in the first case, i.e., as the taste of families for such gender-intrinsic characteristics of children that neither in their extent nor intensity depend on parental outlays. Therefore, the gender bias does not mean that parents want some gender because that will bring higher returns to their investments. Instead, it means that they want some gender because of its predetermined characteristics [4]. And if the gender bias as we understand it was the only determinant of parental birth stopping behavior connected to the gender of children, two relationships for the household outcomes would likely hold. First, parents who desire boys but have a girl or vice versa anticipate having more births in the future and might start saving or work more (Barcellos et al., 2014). Second, parents having children of a preferred gender should spend more on household public goods. That is because their marriage is more stable since with a more preferred child it generates higher surplus (Lundberg, 2005). Therefore, in countries, where firstborns of some gender have on average less siblings (are of preferred gender), parents of firstborns of this gender should work less, save less, and spend more on household public goods. Moreover, if sons directly increase the utility of fathers, then a standard bargaining model of the household predicts a shift of household resources from fathers to mothers. This redistribution could be observable as increased leisure among mothers of sons, or increased consumption of private commodities typically consumed by women (Lundberg and Rose, 2003b).

Turning to the difference in costs of raising sons and daughters, the literature considers two cases [5]. First, it is when sons and daughters have constant, albeit not

[4] Appendix contains more detailed explanation of the difference between the gender bias and the cost difference.

[5] While we test for the difference in costs of children, it is actually the the difference in "prices" of sons and daughters that we are primarily interested in. The price of a child is the commitment of resources required to raise a child of given 'quality'. At the same time, the cost of a child is a measure of the actual amount of resources committed to child-raising (Bradbury, 2004). Thus, the cost of children is deliberately chosen by parents and, in principle, is measurable. In most theoretical models related to the subject, which do not allow of variable quality of children (Dahl and Moretti, 2008; Leung, 1991) , the price of children is constant and equals cost because parents are assumed to pay full life-time prices of children once they are born or the per-period price every

necessarily equal, cost. Assumption of constant costs of children is taken in much if not most of the applied studies on the topic (van Praag and Warnaar, 1997) which frequently calculate so called *normative budgets* [6]. The nominal expenditures or normative budgets, however, do not equal total expenditures on children. The last also include time costs net of the value of children's production. Still, the monetary outlays per se do not fullly reflect the quality of inputs. Another issue is whether parents take into account net flow of future transfers from children (Blacklow, 2002; Adda et al., 2016). Available empirical evidence suggests that parental expectations are important for parental spending (Hao and Yeung, 2015). These assumptions describe a case when parents rely upon some rules of thumb when deciding about outlays on children. These rules of thumb, in turn, are based on perceptions about proper arrangement of living in a given society in a given time (Kornrich and Furstenberg, 2007). Then, to calculate the gender difference in costs of children, studies in the literature use two ways. The first way is based on the adult-good or Rothbarth method of measuring the cost of children. This method, unlike the normative budgets or discretionary equivalence scales (van Praag and Warnaar, 1997), is theoretically plausible (Deaton and Muellbauer, 1986). This method estimates a difference in consumption of private adult good or the leisure time (Bradbury, 2004) between parents having first-born sons and first-born daughters. The second method measures gender difference in costs of children relying upon the subjective scales method (Leyden approach) proposed and substantiated in van Praag and Warnaar (1997).

The second case considered in the literature regarding the difference in raising cost of sons and daughters is when the cost consists of fixed and variable components.

period.

[6]For example, the U.S. Department of Agriculture (USDA) has provided estimates of expenditures on children since 1960. Forensic economists use these figures in wrongful death and birth cases, as well as in child support cases (Lino and Carlson, 2010). The constant cost of children is also assumed in, e.g., Dahl and Moretti (2008); Hazan and Zoabi (2015); Leung (1991); Sienaert (2008); Bojer (2002); Raurich and Seegmuller (2017)

This case is captured by models like those in, e.g., Galor (2011); de la Croix and Doepke (2003); Hazan and Zoabi (2015). In this case either fixed (one-time costs) or variable components (price of human capital) of the child cost could differ. Difference in fixed costs is revealed in parental outlays during the very early years of children (rearing costs). At the same time, the difference in variable component is revealed in difference in availability of parental investment items. Children with lower price of human capital will receive higher outlays and have less siblings due to substitution of quality for quantity (Galor, 2011; Aaronson et al., 2014) [7]. We use a set of home items, which are used as measures of parental investment (Cunha et al., 2010), as proxy variables for parental outlays on children. Parents will buy more of such items in the case they bring more parental utility per unit of expenses for some gender and will have less children after a firstborn of that gender. In our analysis we assume the raising costs to be as in the second case so that it is consistent with economic theory. Thus, if the differential cost explanation is true, parents of a child of the more expensive gender should have fewer children thereafter, spend less on themselves (both parents simultaneously), spend less on adult public goods (due to lower available means to spend) and spend more on children. Moreover, parents of a "more expensive" child should report higher sums needed to make ends meet. However, if the gender bias explanation specified above is correct, they would rather report lower sums. That is because they should spend more on household public goods which exhibit returns to scale in consumption. Restriction on children's age applied in our analysis make sure that children's earnings do not cofound the obtained estimates. That is because we analyze only households in which the oldest child is at most 12 years old which is compulsory schooling age in all European countries.

[7] It could be that either items for some gender are cheaper or produce more of parental utility through children's human capital. One more case is possible when items generate little human capital thus, more of them are bought (i.e., the demand for them is inelastic). But, it is unlikely that this effect would be stronger in countries with more gender-equivalent attitudes.

The two considered causes might actually be in play simultaneously, but our testing points out to the primary cause which is driving the estimates. We expect to find support for the gender bias and no impact of differential costs because cost difference should play a lesser role in European economies (Brockmann, 2001). However, we find evidence that in countries with observed daughter preference it is driven by higher parental expenditure on daughters which is in turn caused by lower price of children's human capital for daughters. Whereas in countries with manifested son preference it is driven by son bias which outweighs the effects of a higher cost of dauhters (which is, however, not as high as in daughter preferring countries). Also, the cross-country correlation between our estimates of the gender preference and the cost difference is stronger than the correlation between our estimates of the gender preference and the conventional measures of gender equality (GGI, GDI, etc.) which arguably approximate the gender bias. All mentioned findings taken together indicate that the gender preference is more strongly determined by the cost difference than by the gender bias. In this case a policy could subsidize cost of human capital for sons from families which are less well off [8].

3 Data and sample statistics

We use a data set from the European Union Survey of Income and Living Conditions (EU-SILC) for years 2004 - 2015. It is a data set collected annually by national statistical offices in cooperation with Eurostat from nationally representative samples and covering the EU-28 and several non-EU countries in 2015. In 2004 only 15 countries were covered by the survey. Our analysis is based on data from four Balkan countries and three Scandinavian countries. The Balkan countries are

[8] A recent study (Scott et al., 2018) finds much lower upward earnings mobility for black men in US than for black women

Bulgaria, Croatia, Slovenia, and the Republic of Serbia [9]. The Scandinavian countries are Denmark, Norway, and Sweden according to the conventional definition of Scandinavian countries [10]. A primary goal of EU-SILC is to collect cross-sectional and longitudinal (using a rotational four-year panel scheme) microdata on income, poverty, social exclusion, and living conditions (Eurostat, 2017). The longitudinal component is not used in our research. The reference population in EU-SILC includes all private households and their current members residing in the territory of the countries at the time of data collection. All household members are surveyed, but only those aged 16 and more are interviewed. The data set for each particular year after 2004 consists of two groups of variables: primary and secondary. Primary variables are collected each year. Secondary variables are collected every five years or less frequently in the so-called ad-hoc modules. A variable may include information either at household or personal level about specific topics. The primary variables convey information on household demographic composition, incomes, living conditions, and labor market activity. The secondary variables used in the current research were collected in years 2009, 2010, and 2013-2015 in ad-hoc modules on material deprivation. These secondary variables contain more in-depth information on material deprivation in the household than annual primary variables. Eurostat calculates cross-sectional household and individual weights to correct for non-random sampling and non-response (Eurostat, 2015) [11].

[9] These are slavic-speaking Balkan countries covered by EU-SILC survey. When we extend the set of considered Balkan countries to include also Greece and Romania, the estimates of the gender preference do not change qualitatively.

[10] These groupings of countries have been frequently used in previous studies. For instance, Estrin and Uvalic (2014) use similar grouping for the Balkan countries and conduct regression analysis on the pooled sample of data from these countries under assumption that regression parameters do not differ between these countries. Similarly, Baranowska-Rataj and Matysiak (2016) and Ragan (2013) use the mentioned grouping of Scandinavian countries. Both studies assume that considered characteristics of economies (model parameters) are similar across Scandinavian countries. In a similar vein, Filmer et al. (2009) pool HNS data into six sub-samples by parts of the world and assume no difference in parameters between countries within groups.

[11] More detailed information on the dataset is available at the following link `http://ec.europa.eu/eurostat/web/microdata/overview`

Table 1: Descriptive statistics - demographics and labour market information.

	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Married couples	
Selected household characteristics	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
Living without	0.114	-0.005	-	-	0.106	0.003	-	-
father	(0.318)	(0.003)	-	-	(0.308)	(0.003)	-	-
Number of children	1.855	0.047	1.872	0.046	1.996	0.004	2.016	0.005
	(1.047)	(0.010)***	(0.996)	(0.010)***	(0.839)	(0.007)	(0.832)	(0.007)
First-born girl	0.481	-	0.484	-	0.487	-	0.487	-
	(0.500)		(0.500)		(0.500)		(0.500)	
Age of mother	26.44	0.035	27.06	0.03	28.91	0.04	29.09	0.07
at first birth [a]	(7.35)	(0.07)	(5.36)	(0.06)	(5.36)	(0.04)	(4.79)	(0.04)
Age of mother	34.68	0.001	35.4	0.0009	37.44	0.002	37.56	0.05
	(7.40)	(0.07)	(6.12)	(0.06)	(6.26)	(0.05)	(5.80)	(0.05)
Mother having	0.178	-0.005	0.195	-0.007	0.363	0.002	0.402	0.004
tertiaty education	(0.382)	(0.003)	(0.396)	(0.004)	(0.481)	(0.004)	(0.490)	(0.004)
Mother employed	0.606	0.000	0.650	-0.003	0.746	-0.001	0.821	0.001
	(0.489)	(0.004)	(0.477)	(0.005)	(0.435)	(0.003)	(0.383)	(0.003)
Mother's weekly	28.100	-0.106	28.738	-0.159	27.985	0.341	28.001	0.340
hours of work	(19.424)	(0.183)	(19.141)	(0.186)	(14.872)	(0.122)**	(14.851)	(0.123)**
Father employed	-	-	0.805	0.004	-	-	0.924	-0.005
	-	-	(0.396)	(0.004)	-	-	(0.264)	(0.002)
Father's weekly	-	-	37.156	0.082	-	-	37.810	-0.165
hours of work	-	-	(16.689)	(0.162)	-	-	(12.762)	(0.106)
Household disposable	20,469.770	265.421	20,982.732	214.079	64,070.609	325.596	65,957.259	450.271
income (euros)	(15,431.683)	(141.036)	(15,550.905)	(150.036)	(57,680.462)	(447.583)	(59,032.599)	(483.734)
Living in urban area	0.137	0.003	0.131	0.002	0.347	0.000	0.341	0.002
	(0.344)	(0.003)	(0.337)	(0.003)	(0.476)	(0.004)	(0.474)	(0.004)
Ownership of	0.767	-0.003	0.763	-0.004	0.920	-0.005	0.929	-0.004
accomodation	(0.423)	(0.004)	(0.425)	(0.004)	(0.271)	(0.002)**	(0.257)	(0.002)**
N of hhds	24,951		22,027		28,352		25,294	

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Note: The statistics were calculated for the entire sample of families with children and for intact families only. Columns one and three provide means and standard deviations while columns two and four provide differences between mean values for girls versus boys. Values in parentheses in even numbered columns correspond to t-test standard errors.

[a] These statistics were calculated only for families in which the mother is younger than 41 and older than 17 and had the first child at the earliest at the age of 16 and children's ages are in the range 0–14.

Two main advantages of this data set are important for answering testing. First, it contains information on age and gender of all adults and their children living in the household. Second, the ad-hoc modules from 2009, 2010 and 2013-2015 contain detailed information on material condition of adults and children in the household. But, there are also two significant drawbacks. First, not all children might be present in the household at the time of the survey for some reason (e.g., because they study or work elsewhere). That is why we cannot be sure that the firstborn (i.e., the oldest) child lives in the household. Second, the information on material conditions of children is available only for all children in the household together and not for every child separately [12]. To correct for the first drawback, we limit our sample for being able to claim with high certainty that the firstborn child is still in the household. Specifically, following other studies in the literature (Dahl and Moretti, 2008; Karbownik and Myck, 2017; Ananat and Michaels, 2008), we limit the analysis to mothers aged between 18 and 40 who had their first child at the earliest at the age of 16. The limit for the age of the oldest child is set at 14 years [13]. Our calculated sex-ratio for firstborns is 1.057, close to the commonly accepted value of 1.06 (Grech et al., 2002) [14]. And to correct for the second drawback, we connect the material condition of children in the household to the gender composition of children (i.e., the share and presence of daughters among children which are instrumented with

[12]For example, an answer to a question: "Do children have books at home suitable for their age?" should be "Yes" if all children have books and "No" if at least one child does not have books.

[13]The sample bias is likely to be very small because the minimal age of leaving school in all European countries is above 16. Other studies(Dahl and Moretti, 2008; Karbownik and Myck, 2017) use the threshold of 12 years. But, Karbownik and Myck (2017) use this threshold since it corresponds to the grouping of expenditure information on clothing. Instead, we need broader range of ages because we aim to control for the age of children (which was not done in other studies). Moreover, Dahl and Moretti (2008) find the 12-year cutoff conservative while Ichino et al. (2011) and (Ananat and Michaels, 2008) use 15-year and 17-year cutoffs respectively. Importantly, the chosen threshold makes sure that children's earnings do not confound our results because this threshold is below the compulsory schooling age in all European countries. At the same time, when we estimate our models on the entire sample, the estimates preserve signs and statistical significance but reduce in size

[14]This fact also suggests that the gender-selective abortion or gender difference early childhood treatment should be to rare for showing up in the data

a dummy for the first child being a girl). Since the gender of children influences the household composition, we limit our analysis, for the most part, to the sample of married couples. The Table 1 contains descriptive statistics for selected household socio-demographic characteristics separately for all families and for cohabiting couples. Table 2 presents descriptive statistics on variables characterising different aspects of the household material condition. We use variables in Table 2 as dependent variables and variables in Table 1 as covariates. Table 2 presents among adult and household material deprivation characteristics also the average frequency of the ten home environment items for children along with girl-boy differences. One can readily see that girls are more likely to have books, have an opportunity to invite friends, or make a celebration. These differences are small, however, and hover around one percent of the standard deviation of the corresponding items. It is less than reported by Xu (2016). The biggest differences between all families and intact families appear to be for food and clothing. Specifically, the girl-boy difference is significant for all families, but disappears for intact families. This could be explained by more limited resources of incomplete families [15]. Otherwise, the intact families do not appear to differ systematically from all families along the considered characteristics. That supports our decision to focus the analysis on intact families.

[15]This result is consistent with the Trivers-Willard hypothesis. But, further exploration of this question is beyond the scope of this study.

Table 2: Availability of selected items in home environment for girls and boys.

	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Married couples	
Dependent variables	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
Household-level material condition characteristics[a]								
Amount of money needed	1,486.629	11.577	1,507.179	8.119	4,725.007	44.569	4,823.201	84.074
to make ends meet	(830.649)	(7.705)	(831.329)	(8.141)	(13,992.615)	(115.330)	(14,112.091)	(122.862)
Ability to make ends meet	0.215	0.004	0.225	0.002	0.776	0.002	0.798	0.006
	(0.411)	(0.004)	(0.418)	(0.004)	(0.417)	(0.003)	(0.402)	(0.003)**
Replacing worn-out	0.278	-0.008	0.290	-0.005	0.888	-0.006	0.905	-0.004
furniture	(0.448)	(0.007)	(0.454)	(0.007)	(0.316)	(0.005)	(0.293)	(0.005)
Adult-specific material condition characteristics[b]								
Ability to spend a small amount of money on oneself (women)	0.522	0.000	0.533	-0.000	0.399	0.017	0.381	0.016
	(0.500)	(0.007)	(0.499)	(0.007)	(0.490)	(0.007)**	(0.486)	(0.007)**
Ability to spend a small amount of money	0.540	0.003	0.573	0.005	0.383	-0.013	0.408	-0.014
on oneself (men)	(0.498)	(0.007)	(0.495)	(0.007)	(0.486)	(0.007)*	(0.492)	(0.007)**

Table 2 (continued)

Dependent variables	Balkan countries: All families: Mean	Balkan countries: All families: Girl-boy difference	Balkan countries: Married couples: Mean	Balkan countries: Married couples: Girl-boy difference	Scandinavian countries: All families: Mean	Scandinavian countries: All families: Girl-boy difference	Scandinavian countries: Married couples: Mean	Scandinavian countries: Married couples: Girl-boy difference
Availability of two pairs of properly fitting shoes (women)	0.615	-0.003	0.627	-0.001	0.437	0.017	0.411	0.015
	(0.487)	(0.007)	(0.484)	(0.007)	(0.496)	(0.007)**	(0.492)	(0.007)**
Availability of two pairs of properly fitting shoes (men)	0.597	-0.000	0.634	0.002	0.408	-0.012	0.435	-0.012
	(0.490)	(0.007)	(0.482)	(0.007)	(0.492)	(0.007)*	(0.496)	(0.007)*
Replace worn-out clothes (women)	0.540	0.003	0.555	0.004	0.415	0.013	0.393	0.011
	(0.498)	(0.007)	(0.497)	(0.007)	(0.493)	(0.007)*	(0.488)	(0.007)
Replace worn-out clothes (men)	0.535	0.002	0.571	0.003	0.396	-0.012	0.422	-0.013
	(0.499)	(0.007)	(0.495)	(0.007)	(0.489)	(0.007)*	(0.494)	(0.007)*
Get together with friends/family at least once a month (women)	0.552	0.004	0.565	0.005	0.429	0.018	0.405	0.017
	(0.497)	(0.007)	(0.496)	(0.007)	(0.495)	(0.007)**	(0.491)	(0.007)**
Get together with								

Table 2 (continued)

	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Married couples	
Dependent variables	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
friends/family at least once a month (men)	0.551	-0.002	0.586	-0.001	0.401	-0.016	0.426	-0.016
	(0.497)	(0.007)	(0.493)	(0.007)	(0.490)	(0.007)**	(0.495)	(0.007)**
Regularly participate in a leisure activity (women)	0.233	-0.006	0.244	-0.006	0.322	0.010	0.307	0.010
	(0.423)	(0.005)	(0.430)	(0.006)	(0.468)	(0.007)	(0.462)	(0.007)
Regularly participate in a leisure activity (men)	0.254	-0.005	0.276	-0.006	0.317	-0.009	0.338	-0.009
	(0.435)	(0.006)	(0.447)	(0.006)	(0.465)	(0.007)	(0.473)	(0.007)
Children's home environment items[d]								
Replacing worn-out clothes	0.822	-0.007	0.843	-0.005	0.986	0.000	0.987	0.001
	(0.382)	(0.007)	(0.363)	(0.007)	(0.118)	(0.003)	(0.113)	(0.003)
Two pairs of properly fitting shoes	0.845	0.006	0.867	0.007	0.983	0.000	0.986	-0.002
	(0.362)	(0.006)	(0.340)	(0.006)	(0.128)	(0.003)	(0.118)	(0.003)

Table 2 (continued)

	Balkan countries				Scandinavian countries			
	All families		Married couples		All families		Married couples	
Dependent variables	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference	Mean	Girl-boy difference
Fresh fruits and vegetables once a day	0.866	-0.010	0.885	-0.006	0.982	-0.003	0.983	-0.003
	(0.341)	(0.006)	(0.319)	(0.006)	(0.134)	(0.003)	(0.127)	(0.003)
One meal with fish, chicken or meat (or vegetarian equivalent) at least once a day	0.842	-0.003	0.862	-0.001	0.988	0.003	0.989	0.002
	(0.365)	(0.006)	(0.345)	(0.006)	(0.108)	(0.002)	(0.103)	(0.002)
Books at home suitable for children's age	0.844	0.006	0.863	0.009	0.983	0.006	0.984	0.005
	(0.363)	(0.006)	(0.344)	(0.006)	(0.131)	(0.003)	(0.126)	(0.003)
Outdoor leisure equipment	0.821	-0.001	0.841	0.004	0.987	-0.002	0.990	-0.003
	(0.383)	(0.007)	(0.366)	(0.007)	(0.112)	(0.002)	(0.102)	(0.002)
Indoor games	0.875	-0.002	0.891	0.000	0.995	-0.000	0.996	-0.001
	(0.331)	(0.006)	(0.312)	(0.006)	(0.072)	(0.001)	(0.066)	(0.001)
Regular leisure activity	0.503	0.010	0.518	0.009	0.776	0.017	0.779	0.019
	(0.500)	(0.009)	(0.500)	(0.009)	(0.417)	(0.008)**	(0.415)	(0.009)**

Table 2 (continued)

Dependent variables	Balkan countries: All families Mean	Balkan countries: All families Girl-boy difference	Balkan countries: Married couples Mean	Balkan countries: Married couples Girl-boy difference	Scandinavian countries: All families Mean	Scandinavian countries: All families Girl-boy difference	Scandinavian countries: Married couples Mean	Scandinavian countries: Married couples Girl-boy difference
Celebrations on special occasions	0.867	-0.002	0.884	0.000	0.981	0.001	0.983	0.002
	(0.339)	(0.006)	(0.320)	(0.006)	(0.137)	(0.003)	(0.129)	(0.003)
Invite friends around to play	0.790	0.002	0.807	0.005	0.959	0.002	0.959	0.002
	(0.408)	(0.007)	(0.395)	(0.007)	(0.198)	(0.004)	(0.198)	(0.004)

Note: The statistics were calculated for the entire sample of families with children and for intact families only. Columns one and three provide means and standard deviations while columns two and four provide differences between mean values for girls versus boys. Values in parentheses in even numbered columns correspond to t test standard errors.

[a] The amount of money needed to make ends meet and the ability to make ends meet are primary variable collected annually while replacing worn-out furniture was collected in ad-hoc modules in years 2009 and 2013-2015.

[b] Adult-specific material condition characteristics were collected in ad-hoc modules in years 2009 and 2013-2015.

[c] This variable and the three next variables were collected in year 2010.

[d] Children's home environment items were collected in ad-hoc modules in years 2009 and 2013-2015.

4 Empirical Analysis

Our paper tests between two alternative explanations of the parental gender preference. Each of the two has different implications for the household economic behavior. The gender bias hypothesis implies that households with a first-born child of the desired gender save less (Barcellos et al., 2014) [16] and spend more on household public goods (Lundberg, 2005). We do not have a direct measure of household savings so we use the capacity to face unexpected financial expenditures as a proxy variable for it. Here we rely on the intuitively appealing assumption that higher savings mean higher capacity to face unexpected expenditures. Regarding the measure of household public goods, we use replacing worn-out furniture. Other measures, like good nutrition and quality of leisure or availability of appliances and cars, are more likely to have direct impact on children's well being and thus might be not invariant to the gender of children. Also, more of the household public goods available should be also reflected in a higher ability to make ends meet and a lower amount of money needed to make ends meet because consumption of household public goods exhibit returns to scale. At the same time, the differential costs hypothesis implies that parents of a child of the preferred gender (i.e., of the more expensive gender so that less of next births follow) work more (leisure substituted for consumption), save less (less of means to save), and spend less on adult public goods (less of means to spend). Parents of more expensive child should report lower ability to make ends meet together with higher sums needed to make ends meet.

One possible way to test our hypotheses is to compare families with different gen-

[16] These authors also mention that in such households mothers finish the maternal leave earlier. The evidence from US, however, suggests that fathers of sons tend to work less. At the same time, many authors find sons to be preferred in US. The descriptive statistics for the pooled EU-SILC sample show that mothers of daughters actually work more while daughters are preferred gender. Nevertheless, a comprehensive testing of this implication for the EU-SILC data is beyond the scope of this paper.

der composition of children. It is the approach taken by Bogan (2013) who explores the relationship between the household stock market participation and the gender of children. Specifically, the author estimates a regression in which the dependent variable is stock or bond ownership while the explanatory variables are dummies for only female and only male children or a proportion of female children in the household. Such estimates, however, cannot be taken as evidence for the causal relationship between the variables in question. That is because the explanatory variable in both specifications, the dummies for the same-gender children and the proportion of daughters, might be decided by households and, thus, be endogenous [17]. Similarly, in the case of our analysis, more daughter-preferring parents could also derive more utility from well-being of children and, thus, tend to create better material conditions for them. To address these concerns we use the gender of the firstborn as the explanatory variable. And our identification strategy is to assume that the gender of the firstborn is randomly determined. Such assumption has been made in other studies which mostly use the gender of firstborns as an instrument for household characteristics. Some of these characteristics are: the bargaining power of women in China (Li and Wu, 2011), the number of children in a family (Dahl and Moretti, 2008), divorce occurrence (Bedard and Deschenes, 2005; Ananat and Michaels, 2008), accommodation area (Dujardin and Goffette-Nagot, 2009) [18]. To test our hypotheses, we proceed in three steps. First, we estimate the gender preference across European countries using the third-parity method. Second, we verify validity of the gender bias explanation by testing its aforementioned implications in daughter-preferring countries and son-preferring countries separately. That is, in countries for which we observe the daughter-preference, parents of the first-born

[17] More daughter-preferring families, for instance, are more likely to have all daughters: they self-select into having all daughters because son-preferring families who have only daughters are more likely to progress to the next parity until they have a son. At the same time, daughter-preferring families could be less risk-averse and, consequently, more inclined to stock market participation.

[18] Appendix contains some additional considerations and reservations about using this instrument.

daughter should be less capable to face unexpected financial expenditures (because they should save less), spend less on themselves, be more likely to replace worn-out furniture, be more able to make ends meet and need less money to make ends meet. Similarly, the same predictions should hold for parents of first-born sons in son-preferring countries. Third, we verify validity of the differential costs explanation by testing its implications similarly in daughter- and son-preferring countries. We do this in two stages. At the first stage we assume the constant costs (prices) of sons and daughters (e.g., Dahl and Moretti, 2008; Jayachandran and Kuziemko, 2011; Leung, 1991). At the second stage we relax this assumption and, instead, assume the cost of children to consist of two components, fixed and variable (e.g., Galor, 2011; Aaronson et al., 2014; de la Croix and Doepke, 2003). In the last case, we find out whether the difference is driven by the fixed or the variable component.

The baseline specification of the regression model takes the following form:

$$y_i = \beta\mathit{(First\ child\ girl)}_i + \boldsymbol{\alpha}\boldsymbol{X}_i + \epsilon_i \tag{1}$$

where y_i stands for either the third parity progression or a children's material conditions indicator for a household i and $\boldsymbol{X}_i$ is a vector of household i socio-demographic and economic characteristics. The *First child girl* indicator takes value 1 if the first-born child was a girl and 0 - if a boy. And ϵ_i is the residual, values of which within a given country can be correlated. A specific set of variables entering $\boldsymbol{X}$ depends on a particular regression equation specification. We use this form at each of the three steps of testing the hypotheses.

To test for the gender preference we put the third parity progression on the left-hand side. Progression to the third parity has been the most widely used indicator in the literature to test for the gender preference. There are two main reasons why it is better to use the parity-three progression rather than the parity-two progression to measure the gender preference. First, it is likely that the desire for the gender-mix of children (to have at least one son and one daughter) coexists with the gender bias towards one gender (Dahl and Moretti, 2008). In that case parents who have bias towards any gender will progress to the parity two independently of the gender of their firstborn. That is why the effect of the gender of the firstborn on the progression to parity two is not likely to be significant. The second reason is that the first-born twins would distort the estimates for the parity two progression. Still, we also report in the results for the second parity progression and for the total number of children on the left-hand side. We choose covariates to put on the right-hand side which are also used in other studies on the subject. These covariates are: gender of the first two children, cubic polynomial in mother's age, squared polynomial in mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area(Dahl and Moretti, 2008; Hank and Kohler, 2000; Haughton and Haughton, 1998; Larsen et al., 1998; Clark, 2000; Basu and De Jong, 2010). We include the higher degree polynomials in the mother's age to take account of conclusions reached by Yamaguchi and Ferguson (1995) who argue that the probability of birth for women is lower at a younger age, then increases, and then again decreases. Such relationship is best fit by the third-degree polynomial in age. Also, we add the family's tenure status in relation to the occupied accommodation along with year and country dummies. We estimate the models with OLS as most of other studies on the subject because this method yields consistent estimates of the coefficient on the gender of the firstborn dummy. Also, the linear probability model may be an especially good choice because right-hand side variables are mostly dummies (of 23 covariates only 7 are continious variables) and the unboundedness problem is less acute in this case (Wooldridge, 2002, p.456). Nevertheless, we also run Probit estimation to check for consistency with the OLS-based results [19]. Since we expect observations not to be iid but correlated within countries, we cluster the standard errors at the country level.

In regard to the testing for differential costs of sons and daughters, we assume that the cost of children consists of two components, the constant (one-time cost) and the variable (outlays on human capital). Researchers commonly use this assumption in models featuring parental investment in children. The fixed component of children's cost represents primarily the time cost of rearing children during infancy. Whereas the variable component represents parental expenditures on children's human capital. Thus, if the previously described analysis reveals that parental outlays on children of some gender are larger, it could have two causes: the larger one-time costs and the lower price of human capital (i.e., of parental discounted utility derived from children's human capital). The mechanism behind the second cause is that of substitution of quality for quantity of children. That is parents spend more on daughters' "quality" and have less children after bearing a daughters. If this explanation is true, daughters in daughter preferring countries should receive more of parental investments. One measure of parental investments used in the literature [20] is availability of conditions and items at home which are necessary for children's normal development (Cunha et al., 2010; Todd and Wolpin, 2007; Juhn et al., 2015) [21]. The expected effects of the first-born daughter are systematically presented in the Table A7. We use EU-SILC data on availability of such items in households in 2009 and 2014 to test if daughters tend to have better material conditions in daughter-preferring countries and sons, respectively, in son preferring countries. Moreover, under such assumption, parents having a child of the more expensive gender, in addition to having lower progression ratio, should also have lower expenditures on

[19]Probit estimates actually correspond to OLS estimates in terms of impact direction and statistical significance.

[20]The most common measure is the years of schooling conditional on household income.

[21]These variables are described in more detail in the Appendix

private consumption and household public goods, be less able to face unexpected financial expenditures, be less able to make ends meet, and need more money to make ends meet. The ability to make ends meet is measured by a binary variable taking value 1 when a give household is able to make ends meet. The aforementioned predictions follow from the fact that they have less financial means left after making outlays on children than parents having a child of the cheaper gender. The method of measuring the cost of children through comparing the amount of money needed to make ends meet reported by families having children of different gender was proposed and used by van Praag and Warnaar (1997).

5 Results

5.1 Estimates of the preference for the gender of children at birth

First, we present results of testing for the gender preference in two groups of countries. The Table 3 presents coefficients on the gender of the firstborn for different specifications of the dependent variable in the Equation 1 estimated on data from the Balkan countries. These results resemble the results obtained by Dahl and Moretti (2008) for US. The first column indicates that families where the first child is a girl end up having more children than families where the first child is a boy, although the difference is not significant. In line with the expectations discussed above, the impact of the gender of the firstborn on progression to parity two in cloumn (2) is much less statistically significant than the impact on progression to parity three and has much lower percent effect. Numbers in column (3) show the probability of having three or more children is 1.3 percentage points higher when the first child is a girl which is an order of magnitude higher than the result obtained by Dahl and Moretti (2008) for US. In other words, first-born girl families are 17% more likely to have three or more children compared to first-born boy families. We also find significant positive effects for the probability of four or more and five or more children when the first-born child is a girl. The positive effect of the first-born daughter on progression to parity three has also been found by Filmer et al. (2009) for Central Asia, South Asia, Middle East, and North Africa. It is this result which is most commonly interpreted in the literature as manifestation of the son preference.

Table 3: The firstborn-child gender and fertility in Balkans.

	Breakdown by number of children				
	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children
First-born child being a girl	0.030	-0.001	0.013	0.011	.003
	(0.010)***	(0.008)	(0.005)***	(0.002)***	(0.001)***
Controls	Yes	Yes	Yes	Yes	Yes
First boy baseline	1.57	0.483	0.077	0.011	0.002
Percent effect	0.019	-0.002	0.17	0.18	0.50
R-sq	0.26	0.39	0.13	.04	.02
Observations	19,807	-	-	-	-

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: S.E. are given in parentheses and are clustered at the country level. Estimates are based on the 2004-2015 EU-SILC samples for Bulgaria, Croatia, Serbian Republic, and Slovenia. The sample consists of households formed by a one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had the first child at the earliest at the age of 16 and children's ages are in the range 0–12. The estimation method used is weighted OLS with probability weights reflecting non-random sampling within and between countries. The table presents estimated effects of the firstborn being a daughter compared the baseline case of the firstborn being a son. The percent effect is a ratio of the estimated OLS coefficient on the firstborn's gender dummy to the baseline value of thee dependent variable. The dependent variables are the total number of children and a set of binary indicators for specific numbers of children. The control variables besides the gender of the firstborn are: the dummy for a first-born daughter, gender of the first two children, cubic polynomial in mother's age, squared polynomial in mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies.

The Table 4 presents estimates analogous to those in the Table 3 for Scandinavian countries. These results are notably different from the results for Balkan countries. First, the impact of the first-born daughter on progression to parity three in column

Table 4: The firstborn-child gender and fertility in Scandinavia.

	Breakdown by number of children				
	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children
First-born child being a girl	-0.009	0.002	-0.013	0.002	0.0002
	(0.010)	(0.006)	(0.005)***	(0.002)	(0.0002)
Controls	Yes	Yes	Yes	Yes	Yes
First boy baseline	1.82	0.64	0.16	0.02	0.003
Percent effect	0.005	0.003	0.08	0.1	0.07
R-sq	0.29	0.38	0.22	0.05	0.01
Observations	25,227	-	-	-	-

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: Estimates are based on the 2004-2015 EU-SILC samples for Denmark, Norway, and Sweden. For details about sampling and estimates presentation, see the notes under the Table 3

(3) is negative and statistically significant. Despite it has a similar absolute value, its percent effect is one half of that for Balkans because a larger share of Scandinavian families progresses to the parity three. Second, impacts of the first-born daughter on the total number of children and on progression to other parities have small absolute magnitudes and are not statistically significant. The results on the parity three progression in column (3) are in line with those obtained by Andersson et al. (2006) for each of the Scandinavian countries separately. These results alone suggest that the gender bias is likely to be not the only mechanism behind these results because in that case the results should also be similar for progressions to higher parities.

In Appendix, we also explore the gender preferences across EU countries. Our results are broadly consistent with those obtained in the literature before (Hank and Kohler, 2000). We also attempt to evaluate how our results would differ if there was no family disruptions caused by the gender of children which is frequently reported in the literature (see, e.g., Lundberg (2005) for review). The estimates obtained for that counterfactual scenario, however, do not differ qualitatively and do not differ much quantitatively from those reported here. Absence of rank correlations between the country-level impacts of the firstborn's gender on progression to the parity two and to the parity three suggests different driving causes behind these impacts as we expected above.

5.2 Testing between the gender bias and the differential cost explanations

First, we test implications of the two explanations for the household-level outcomes. The gender bias explanation implies two patterns in the household-level allocations. First, expenditures on household public goods should be higher when the firstborn is of the preferred gender (Lundberg, 2005). Specifically, if a son increases marital surplus more than a daughter, then the birth of a son reduces the probability of divorce, and also increases the incentive of partners to invest further in the marriage, or the family as a whole (Lundberg and Rose, 2003b). Second, saving in such case should be lower because parents anticipate less births in future (Barcellos et al., 2014). To test the first implication, we estimate impact of the first-born daughter on frequency of replacing furniture in the household. Lundberg and Rose (2003b) consider furniture an important household public good along with automobiles and housing conditions as proxied by housing expenditures. The spending on automobiles and housing, however, can be directly influenced by the gender of children. As Lundberg and Rose (2003b) note, observed differences in housing spending would rely on the need for space to accommodate the size and activity of sons, or on a desire for a higher quality neighborhood to reduce the probability of risky behavior by boys. Concerning the automobiles, having one might make more sense when a couple has sons for whose socialization access to automobiles can often be more important. Meanwhile, expenditures on furniture do not appear to be directly influenced by the gender of children. Column (1) of Table 5 contains estimates of the firstborn's gender impact on replacement of worn-out furniture in the household.

The negative and statistically significant estimate for Balkan countries confirms the prediction following from the son bias explanation of the observed gender preference. To support the daughter bias explanation for Scandinavian countries, the estimate needs to be positive which is not the case. Regarding the prediction that savings should be lower in families with a firstborn of the preferred gender, we test them by estimating the impact of the firstborn's gender on the ability to face unexpected expenditures. We assume that households with higher savings are more likely to respond positively to this question, i.e. we use this indicator as a proxy variable for household savings. Thus, if the gender bias explanation is true, the estimate should be positive in Balkan countries and negative in Scandinavian countries. The obtained estimates in column (2), however, are small in magnitude and not statistically significant. For the case of Balkan countries this result could be reconciled with the son preference by the fact that common savings are also a household public good and respond positively to arrival of a child of the preferred gender countering the negative effect of reduction in expected number of children.

The higher expenditure on household public goods is also consistent with the fathers' comparative advantage in raising sons. This is the so called "technology" explanation according to Dahl and Moretti (2008). The gender bias explanation and technology explanation have different implications for consumption patterns of fathers and mothers. The gender bias explanation suggests lower consumption of mothers of daughters while the technology explanation implies it to be higher. Specifically, if sons directly increase the utility of fathers, then a standard bargaining model of the household predicts a shift of household resources from fathers to mothers. This redistribution could be observable as lower consumption of private commodities by mothers of daughters. The negative impact of the mother's ability to spend on oneself in Balkan countries in column (3) of Table 6 is in line with the gender bias explanation. In addition, two more facts hold for intra-household

Table 5: Impact of the first-born girl on availability of household public goods across countries grouped by the observed gender preference

Countries	(1) Replacing worn-out furniture	(2) Capacity to face unexpected expenditures	(3) Ability to make ends meet	(4) Lowest monthly income to make ends meet	(5) Availability of home items
Balkan	-0.020	0.0019	0.008	-0.671	0.017
	(0.011)*	(0.007)	(0.006)	(9.848)	(.015)
Scandinavian	-0.006	0.005	0.005	152.7	0.035
	(0.007)	(0.005)	(0.004)	(142.2)	(0.018)**

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: The standard errors of estimates on sub-samples for Balkan and Scandinavian countries are clustered at the country level. Estimates in the columns (2), (6), and (7) are based on the 2009 and 2013-2015 EU-SILC ad-hoc modules while estimates in remaining columns are based on the 2004-2015 EU-SILC primary modules. The samples consists of households formed by a one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had the first child at the earliest at the age of 16 and children's ages are in the range 0–12. The estimation method used is OLS. Dependent variables for columns (1) and (3)-(7) are binary indicators taking value 1 when a household has the indicated condition and value 0 otherwise. The table contains estimated coefficients on the dummy variable for the first-born child being a girl. Other control variables are: household disposable income, weekly work-hours of partners, dummies for parental employment status (full-time employment, self-employment, part-time employment), tenure status, living in urban area, educational attainment of partners, year dummies, and country dummies.

Table 6: Impact of the first-born girl on employment consumption of mothers and fathers in Balkans

	(1) Being employed	(2) Weekly hours of work	(3) Ability to spend on oneself	(4) Two pairs of shoes	(5) Replacing clothes	(6) Get together with friends	(7) Regualr leisure activity
Mothers	-0.011	-0.369	-0.0233	-0.007	-0.007	0.006	0.032
	(0.006)*	(0.265)	(0.0117)**	(0.01)	(0.01)	(0.01)	(0.024)
Fathers	-0.006	-0.328	0.004	-0.002	-0.003	0.002	0.049
	(0.005)	(0.228)	0.011	(0.01)	(0.01)	(0.01)	(0.024)**

Notes:

The standard errors of estimates on the sub-sample for Balkan countries are clustered at the country level. For details on sampling and estimation see the note under Table **??**.

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

allocations in Balkan countries. First, mothers of daughters are less likely to be employed. Second, fathers of daughters report more time spent on leisure. The first fact with could be explained by self-selection into unemployment of mothers who

Table 7: Impact of the first-born girl on employment and consumption of mothers and fathers in Scandinavia

	(1) Being employed	(2) Weekly hours of work	(3) Ability to spend on oneself	(4) Two pairs of shoes	(5) Replacing clothes	(6) Get together with friends	(7) Regualr leisure activity
Mothers	0.005	0.439	-0.002	0.005	0.001	0.013	-0.060
	(0.005)	(0.185)**	(0.008)	(0.006)	(0.007)	(0.007)*	(0.031) **
Fathers	-0.007	-0.357	0.004	(0.004)	0.005	0.0003	-0.032
	(0.003)	(0.156)**	(0.008)	(0.006)	(0.007)	(0.007)	(0.030)

Notes: The standard errors of estimates on the sub-sample for Scandinavian countries are clustered at the country level. For details on sampling and estimation see the note under Table **??**.
* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

have especially strong comparative advantage in raising daughters (otherwise, the first-born daughters would also have negative impact on intensive margin of mothers' employment). Still, such self-selection of mothers into employment would not undermine our results because the "technology" explanation implies lower progression to parity three in the case if fathers had sufficiently high comparative advantage in raising sons and there were sufficiently wide wage gap in favor of men (Gugl and Welling, 2012). Despite there exists a wide gender wage gap, our estimates do not support existence of a sizable comparative advantage of fathers in raising sons in Balkans which would be evident from lower hours of work and higher personal consumption reported by fathers having first-born sons as explained above. Finally, the fact that fathers have more leisure could be explained by longer hours of housework done by daughters [22]. Thus, the obtained results are consistent with the gender bias explanation for Balkan countries.

Next, we turn to examining implications of the gender bias explanation for Scandinavian countries. First, at the household level there is no firstborn's gender effect

[22] This is true for 2010 ad-hoc sample form Romania and Bulgaria. The question on hours of housework was included in 2010 EU-SILC ad-hoc module optionally by national statistical agencies which is why this data is available for 10 EU countries only.

neither on furniture replacement nor on ability to face unexpected expenditures as estimates in the first two columns of the Table 5 indicate. Moreover, estimates of the firstborn's gender impact on the parental consumption in Table 7 do not differ between fathers and mothers like unconditional means in the Table 3 which would be in line with working of the parental comparative advantage [23]. That is because mothers of sons should redirect through intra-household bargaining process household resources to fathers for keeping them in the family due to their important role in raising sons (Lundberg, 2005). At the same time, estimates of the impacts on ability of mothers to meet with friends and family and on ability to have a regular leisure activity do not contradict the gender bias explanation per se. The estimated impacts on fathers' consumption, however, should be positive according to the gender bias explanation which is not the case. Still, fathers of daughters work less hours. Nevertheless, this is not reflected in larger amount of leisure time enjoyed by them. Moreover, fewer hours worked by fathers of daughters are not likely to drive the observed daughter preference because similar effects were found for US and West Germany (Lundberg and Rose, 2002; Choi et al., 2008) which exhibit son preference (Dahl and Moretti, 2008; Hank and Kohler, 2000). Therefore, the data does not support the gender bias explanation for Scandinavian countries.

Concerning the differential cost hypothesis, it is not confirmed by household-level estimates for Balkan countries. There are no statistically significant results in the last three columns of the Table 3 for Balkan countries. Moreover, if expenditures on sons were higher explaining the lower progression after the first-born son, parents of daughters would have more resources to spend on themselves which contradicts the negative impact of the first-born daughter on private expenditure of mothers in

[23]Despite the explanation of the gender preference through parental comparative advantage does not work in the case of Scandinavian countries, we still consider it as a possible mechanism behind intra-household allocation. It can not be the main driving cause for the observed gender preference in Scandinavia because in that case the gender wage gap should be in favor of women (Gugl and Welling, 2012) which is not the case in Scandinavia.

column (3) of the Table 5. At the same time, our results are in line with presence of higher outlays on daughters in Scandinavian countries. Households with first-born daughters are more likely to have available a whole set of ten important children consumption items. Still, neither ability to make ends meet nor the minimum amount of money to make ends meet depend on the gender of the firstborn. Nevertheless, for the top income decile the minimum amount of money needed to make ends meet is larger for families with the first-born daughter. The argument why this should be true is developed in the Appendix (the Figure A4 illustrates this idea). Turning to estimates of the firstborn's gender impact on parental consumption, mothers of daughters appear to more frequently forgo the regular leisure activity and substitute it with apparently less costly socialization through meeting with friends and family. Moreover, more hours worked by mothers of daughters suggest that they are willing to substitute leisure for outlays on daughters. At the same time, fathers of daughters tend to work less than fathers of sons. When Lundberg and Rose (2002) reported similar effect for fathers from US, they offered an explanation based on the son bias idea but did not formally test it. Our testing, however, does not support the son bias explanation. At the same time, Cools et al. (2015) report based on Norwegian data that paternal leave has more pronounced positive effects for daughters than sons. That could be a reason why fathers in Scandinavian countries might be motivated to substitute time spent on work for time spent on children (rather than leisure). Examining data from detailed time-use surveys could split more light on this issue. All in all, the differential expenses explanation of daughter preference in Scandinavian countries is not rejected by the data.

In Appendix we also analyze cross-country relationships between the gender preference, the gender gap in parental investment, and conventional measures of gender equality. These results appear to reinforce our previous points.

6 Conclusion

We find evidence that the gender preference in different countries is caused by different reasons. In Balkan countries the observed son preference is likely driven by the bias towards sons that plays the major role. In Scandinavian countries the observed daughter-preference is likely driven by lower price of daughters quality (which incorporates the specific personal characteristics and their usefulness for parents). To measure precisely the effect of the gender difference in the cost of children we would need to observe its random variation [24]. The fact that the evidence for the lower price of human capital for girls is most pronounced in more gender-equal societies is in line with trend of institutional change in modern societies in favor of women (Roberts and Baumeister, 2011). If it is not compensated by policy which reduces the price of human capital for sons in less well-off families, the consequences mentioned in Edlund (1999) and Seidl (1995) might become more likely to occur.

[24] For example, Miaari and Sauer (2011) mention about different rules of employment for men and women applied by the government in Israel to the inhabitants of the disputable territories. One could think about changes in these rules as a source of random variation in the net cost of children of different gender.

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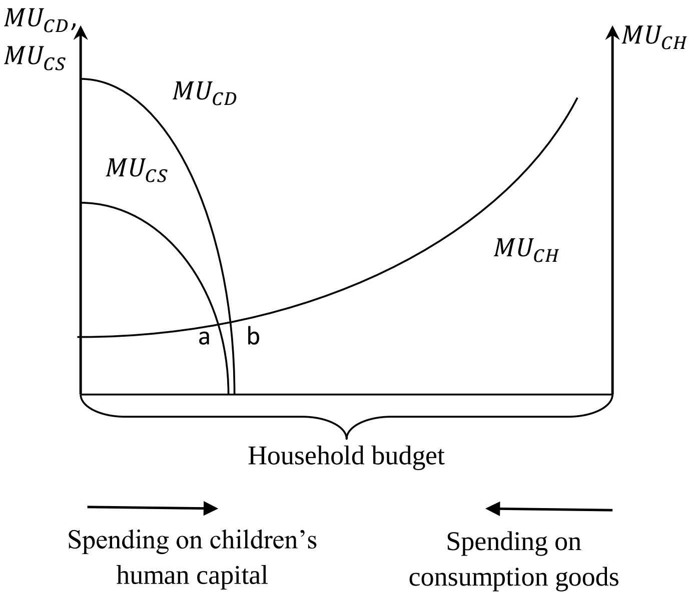
Appendix

The distinction between the gender bias and differential costs concepts

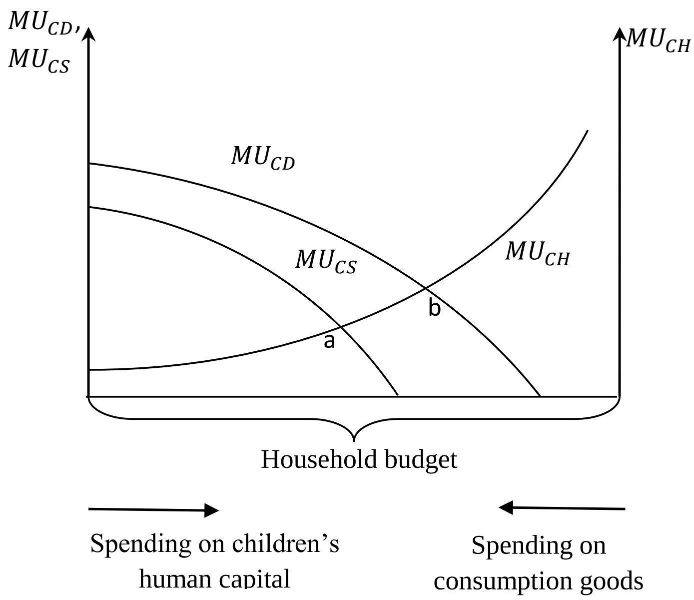
In the literature, there is neither a clear-cut definition of the gender bias nor a conventional term to label it. In some cases, the gender bias is readily recognizable. For example, Arnold et al. (1998) assert that some Indian parents prefer sons for reasons connected with religious beliefs and kinship descent whereas Jacobsen et al. (1999) argue that women's need of companionship leads to the daughter preference in Denmark. Such children's characteristics, continuing the family name or providing the same-gender companionship to parents, are intrinsically pertinent to the gender of children and their utility does not directly depend on the parental outlays on children. Preferences for such characteristics are captured by the first part of the aforementioned definition of Lundberg (2005) because a son has a greater marginal value in the first case and a daughter - in the second. And this understanding is consistent with other previously provided definitions. In other situations, the gender bias is less recognizable. One possible example could be a case man who might want a son because he can become a player in his favorite soccer team. But, he cannot do much beyond encouraging him or taking to a local soccer academy. Had he a daughter, he would likely have done not much less for her physical development. Similarly, parents might want a daughter because she can become a soprano singer. These examples are captured by the second part of the aforementioned definition. That is the father values son's soccer skills more than daughter's in the first example because they increase chances of him becoming a player in a father's favorite team. While in the second example parents value daughter's singing skills more than son's (because the son's soprano will eventually disappear). In both cases, parents would not need to invest much provided the children have sufficient aptitude (taking to a soccer academy or a music school). The education will most likely be done in boarding schools or similar establishments and paid (fully or partially) by parties external to the family (frequently in form of a stipend provided by government or non-profit organizations). A common feature of all previous examples is absence of close relationship between the parental investment of time and market goods on the one side and the child quality (desired characteristics) on the other side beyond some certain relatively low level of investment.

For example, parents want someone from their close surrounding to know a foreign language. The obvious way to proceed is to make a child know that language. On average, it would be cheaper with a daughter because girls are known to be much better in picking up foreign languages (Burman et al., 2008). And the more parents invest in a child's language learning the better is the result (hours with tutors, educational trips abroad, etc.). Keeping other things equal, in such case parents are likely to invest significantly more in the daughter's language learning. Similarly, parents might want their child to earn as a photo model and that is why they attach significant value to children's height. In this case daughter's height is valued much more because women photo models earn four times more than male photo models (Frank, 2008), a rare example of the reverse salary gender gap. Moreover, the children's height depends to some extent on parental investment (nutrition, material living conditions, etc.). Thus, *ceteris paribus*, in this case parents would likely invest more in daughters. The fact, that the sons' height could be more responsive to parental investment (i.e., they are "more productive" in producing the "children's height") does not make a difference because daughters' height ultimately brings much higher utility. We treat such situation as difference in raising costs and discuss it below.

Cases of the gender bias and the differential costs are depicted graphically in Figure A1.



(a) Gender bias



(b) Differential cost

Figure A1: Distinction between cases of the gender bias and the differential costs

Notes: The graphs show marginal parental utilities of expenditures on children's human capital, MU_{CS} and MU_{CD}, together with accompanying marginal utility of household consumption expenditures, MU_{CH}. The underlying unitary household model is assumed. On the horizontal axis there is expenditure on the child's human capital. On the vertical axis is the household marginal utility. The marginal utility of the household consumption is increasing because expenditures on household consumption decrease along the horizontal axis with increase in expenditures on children's human capital. On the left graph marginal utilities of expenditures on children's human capital plummet quickly and parental investments in children's human capital are both low and do not differ significantly between genders. At the same time, the difference in parental utility derived from children of different gender is significant. We assume that this is a case of the gender bias. On the right graph, the marginal utility of investment in a child of some gender is notably larger along a broad range of possible investment volumes. The optimal volumes of investment differ much between children of different genders. We assume that this is a case of the differential cost.

Considerations about using the gender of the firstborn as the instrumental variable

Some authors claim that the gender of the firstborn is not random. For example, Norberg (2004) reports that children who were conceived when their mother was living with a partner were 14 per cent more likely to be boys than siblings conceived when the parents were living apart. This finding aligns with the falling sex ratio in a set of industrialized countries (Davis et al., 1998). One possible explanation for these findings is the evolutionary advantage of species that can adjust the sex ratio of offspring in response to changes in conditions affecting the relative reproductive success of males and females (Trivers and Willard, 1973). And, actually, the wealthiest individuals in societies tend to have sons born more frequently (Cameron and Dalerum, 2009). To address these concerns we repeat our analysis on the sample of partners cohabiting at the time when the firstborn arrived, control for the country fixed-effects, and repeat the analysis after dropping top 1% of wealthiest households in each country from the sample [25].

At the same time, the gender of the firstborn might impact marital stability (Lundberg and Rose, 2003a; Mammen, 2008; Lundberg et al., 2007), family size (Hank and Kohler, 2000; Angrist and Evans, 1998), parental time allocation (Lundberg and Rose, 2002; Lindström, 2013; Choi et al., 2008). This makes "the exclusion restrictions a priori unpersuasive" (Lundberg, 2005). To solve this problem, we focus our analysis on the sample of intact families, instrument the number of children with twin-births, and argue that the impact of the gender of the firstborn on parental employment does not notably alter our estimates.

The documented impact of the gender of firstborns on parental employment differs across countries. For example, a first-born son increases men's workours in US by 3% of mean compared a first-born daughter (Lundberg and Rose, 2002) (on the other hand, Pabilonia and Ward-Batts (2007) find $\frac{1}{3}$ of the effect which is not statistically significant). Even larger effect - almost 5% of mean annual men's workhours - was found for West Germany (Choi et al., 2008). Meanwhile, Ichino et al. (2011) find a negative impact of the firstborn son on mother's working hours and employment in US, UK, and Italy. Still, it is smaller than the previously mentioned effect for fathers and hovers across countries around 1% of the mean. In addition, Lindström (2013) finds that a firstborn son increased fathers' parental leave by 0.6 days (1.5 %) and decreased mothers' leave by a similar amount. In our analysis, we do find

[25] One study (Kanazawa, 2007) reports that physically more attractive parents are significantly more likely to have a daughter. We, however, are not aware of other studies confirming this finding.

the effect of the gender of the firstborn on employment status of mothers. But, we don't find the effect on mothers' work hours and neither on fathers' employment status nor work hours. The negative effect of the first-born son on mother's employment equals approximately 1% of the of the mean women's employment. This is in line with the previously reported estimates from the literature. However, when we multiply this effect on employment on the coefficient on the employment itself, the final effect on the variable of interest is by an order of magnitude smaller than the direct effect of the first-born gender variable. That is why, following Karbownik and Myck (2017) we believe that the impact on employment does not undermine our estimates of interest and proceed keeping the employment status and workload of parents as covariates.

A description of the material deprivation measures

Turning to the EU-SILC ad-hoc modules on material deprivation from years 2009 and 2014, they each contain thirteen questions on availability of children's items and amenities (the module from year 2009 contains questions on 22 items, but the recent module was reduced). Each of these variables indicates presence of a specific item or amenity. These variables are: replace worn-out clothes, two pairs of properly fitting shoes, fresh fruit and vegetables once a day, one meal with fish, chicken or meat (or vegetarian equivalent) at least once a day, books at home suitable for children's age, outdoor leisure equipment, indoor games, regular leisure activity, celebrations on special occasions, invite friends round to play and eat from time to time, participate in school trips and school events that cost money, suitable place to study or do homework, go on holiday away from home at least 1 week per year. We primarily use in our analysis only the first ten questions because they are available for nearly all children in the sample while the last three are available only for school-age children. These questions do not completely correspond to the questions from other surveys on material conditions of children that were analyzed in the literature (for instance, NLSY79-CS HOME-SF module (Cunha et al., 2010; Todd and Wolpin, 2007; Juhn et al., 2015) or PISA-2000 Xu (2016)). Those surveys are more extensive. Instead, the considered ten questions largely overlap with the resources-spent and the-time-with-child subcomponents defined by Juhn et al. (2015) based on the NLSY79 survey. For instance, all questions in resources-spent and some questions from the-time-with-child subcomponents of Juhn et al. (2015) are contained in EU-SILC ad-hoc modules from years 2009 and 2014. All in all, the considered EU-SILC ad-hoc modules could be seen as extended versions of the two subcomponents mentioned just above. And since elements in these two subcomponents were highly correlated with (Bradley and Caldwell, 1980, 1981, 1984) and strongly influencing (Cunha et al., 2010) childrens' development, the raw score of the considered EU-SILC ad-hoc modules should also be correlated with and having impact at children's development. Also, the responses from PISA-2000 survey analyzed by Xu (2016) contain more detailed information corresponding to EU-SILC questions on participating in regular leisure activity, availability of a suitable place to study, and having books at home. And the author argues that items asked about in those questions are important for children's adult outcomes and supports the point by referring to multiple related studies. To test for the gender-gap in the children's material conditions at home, we use five alternative dependent variables in equation 1 for measuring the material conditions. The first one is a sum of ten binary indicators of presence of the ten considered material conditions listed in the previous section. This sum corresponds to the so-called HOME index used in the literature. One problem with this variable is susceptibility to monotonous transformations, the scaling problem (Bond and Lang, 2013) Another problem is that all items under such construction of the dependent variable are assigned equal weights in summation (and, thus, those which have larger variance contribute more to the estimated effect). We try to overcome these problems by constructing four other measures of the material conditions. First, we conduct the principal component analysis (PCA). And the first principal component obtained from this analysis is used as an alternative dependent variable. In this way we follow Cools and Patacchini (2017) who also construct a measure for material conditions of children albeit based on a different data set, using different indicators, and addressing a different research question. The rationale behind such method is elaborated, e.g., by McKenzie (2005). He applies this method to measuring household wealth inequality based on responses about availability of different items. Importantly, he shows invariance of this measure across linear transformations. Also, we use ordered probit and Poisson models with the raw sum of ten indicators as the dependent variable. In this case, however, we assume that households acquire the most necessary children's items first. And the probit and the Poisson regressions measure probabilities of acquiring the next most necessary item. Finally, the frequency histogram of the raw sum of indicators (Figure A1) shows that around one-half of the households possess all ten items. Therefore, we introduce one more binary alternative dependent variable. It takes value 1 for households which possess all specified items and value 0 for other households. This specification of the dependent variable is the most intuitively appealing to us. And we rely upon it in the main analysis. Still, under all specifications of the dependent variable the results of the analysis are qualitatively similar and estimated coefficients of primary interest are statistically significant.

Cross-country comparison of gender preference and parental investment

Table A3 displays the results of estimating the gender preference by country. The geographical pattern of the gender preference at birth is depicted on Figure A6. Our results are broadly consistent with those obtained in the literature before. Like, Hank and Kohler (2000) we find the daughter preference for Portugal and Lithuania and the son preference for Italy and France. Also, like Andersson et al. (2006) we find the daughter preference for Norway. But, they also find the daughter preference in Sweden while we do not [26]. We also attempt to evaluate how our results would differ if there was no family disruptions caused by the gender of children which is frequently reported in the literature (see, e.g., Lundberg (2005) for review). The results are presented in Table A3: son preference becomes statistically significant in Slovenia and stops being statistically significant in Croatia. But, the estimates obtained after including Slovenia in and excluding Croatia from son preferring countries do not differ qualitatively and do not differ much quantitatively from those reported here. The rank correlations between the country-level impacts of the firstborn's gender on the selected household fertility outcomes are presented in Table 4. Absence of a strong correlation between estimated impacts on progression to the parity two and to the parity three suggests different driving causes behind these impacts as we expected above.

Correlation between the second-parity coefficients and the third parity coefficients is quite low (Table A1). But, two measures of the same variable should be correlated. Still, the last two sets of coefficients are strongly correlated with coefficients for the total number of children. This fact might spur a question if it is proper to use the third parity progression for measuring the gender preference, a frequent practice in the literature.

To rationalize the obtained estimates, we plot the coefficients against several existing measures of the gender inequality. As Figure A2 shows, the obtained estimates do not exhibit a strong relationship with those measures. Only the coefficients from the third-parity equation exhibit a negative relationship with our gender equality score based on Eurobarometer data and with the proportion of households reporting the balanced decision-making. At the same time, coefficients for neither the

[26] Still, our estimates are correlated with (ρ=0.6) and statistically significantly predict the estimates by Hank and Kohler (2000)

total-number nor the second-parity equations exhibit any such relationship. This fact suggests once more that the second parity progression and the third parity progression actually measure different kinds of preferences. And that is why we use the third-parity progression results in Figure 1 and beyond.

In addition, the fact that parents tend to invest more in daughters as measured by presence of the considered home items [27] hold for the pooled EU-SILC sample. To test for the gender gap in parental investment we estimate Equation 1 with several alternative measures of children's material conditions on the LHS. We primarily focus on the specification with the binary home indicator (the dummy variable for availability of all 10 considered items at home) on the LHS. The Table **??** displays estimates for this specification on a pooled sample. The results in Table **??** suggest that the daughters on average receive more of parental investments in terms of home environment conditions. For example, the number in column 1 means that in families with first-born girls children are 1.5% more likely to have all 10 considered items available. This estimate is robust to using alternative sets of covariates as could be seen from the rest of the Table **??**. Still, this effect is not large, remaining between 1,7% and 2% of the standard deviation of the binary home indicator. But, results of such scale are typical for the literature on the gender effects as we mentioned above. Meanwhile, the gender preference pattern established before still holds for the sub-sample of households from the highest income decile. These results might suggest that the society as a whole is attaching increasingly positive significance to a female child. This idea has already appeared in, e.g., Brockmann (2001); Andersson et al. (2006). The daughter assumes both the role of a breadwinner and that of a caregiver [28]. As Brockmann (2001) puts it, "in the future, the average girl may well wish to become the mother of a one-daughter family."

[27] Availability of these indicators has been frequenly used in the literature as a measurement of parental investment. More detailed discussion is presented in Section 3

[28] In this regard some authors speak about the "boy crisis" (Husain and Millimet, 2009; Sadowski, 2010).

Like in the case with estimates of the preference for the gender of children at birth, we relate the estimates of the gender gap in parental treatment to specific country-level measures of gender inequality. The gender of the firstborn impact on material conditions exhibits much stronger relationship with the conventional measures of gender inequality than the impact on parity progression. The Figure A3 displays the three strongest relationships. Most importantly, there is a strong relationship with the Global Gender Gap (GGG) score calculated by the World Economic Forum (we used the most recent 2016 data). This index is also strongly related to the gender gap in PISA math achievement (Guiso et al., 2008).

But, Xu (2016) did not find any strong relationship between the gender gap in home environment measure (similar to ours) and the GGG. He, however, measured the gender gap by difference in the unconditional mean between genders. And we explained before that our measure is preferable to the one used in Xu (2016). Therefore, the gender gap in children's material conditions more closely corresponds to conventional gender-inequality measures than the gender gap in the number of younger siblings [29]. Nevertheless, the latter is commonly used as a measure of the gender preference for children.

[29]Similar and statistically significant relationship also holds between the first-daughter coefficient in the material-conditions regression and two other indexes, the GDI (it highly correlates with the GGG) and the SIGI (but it is available only for seven countries from our sample).

Table A1: Coefficients corrected for the selection bias

Cntrs.	Coefs.	Cntrs.	Coefs.	Cntrs.	Coefs.	Cntrs.	Coefs.
AT	0.006	EE	-0.0007	IS	-0.003	PL	-0.003
BE	0.0003	EL	-0.006	IT	0.011***	PT	-0.017***
BG	0.0217***	ES	-0.001	LT	-0.006	RO	0.024***
CH	0.002	FI	0.004	LU	0.003	RS	0.029**
CY	-0.016*	FR	0.007	LV	-0.002	SE	0.010
CZ	0.002	HR	0.027*	MT	-0.010	SI	0.012**
DE	0.006	HU	-0.008*	NL	-0.004	SK	0.010
DK	-0.017**	IE	0.007	NO	-0.018**	UK	0.0007

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: The estimates contained in this table do not differ from those in the third column of the Table A5 except in the sample characteristics and omission of father-related control variables (which have little explanatory power). The sample includes also incomplete families with simulated number of children assuming that those divorced because of the gender of children are characterized by bias towards that gender and do not stop childbearing until they have a desired child.

Table A2: Effects of the gender of the firstborn child in the household on the selected measures of fertility

Explanatory var-s	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children
First child a girl	-0.0050**	-0.0073***	0.0011	0.0004	0.0005*
	(0.0025)	(0.0017)	(0.0012)	(0.0006)	(0.0003)
Controls	Yes	Yes	Yes	Yes	Yes
First boy baseline	1.54	.406	.106	.0248	.00462
Percent effect	-.00323	-.0179	.0102	.018	.109
R-sq	.27	.235	.137	.0491	.0163
Observations	265,507	265,507	265,507	265,507	265,507

*

$p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: S.E. are given in parentheses and are clustered at the country level. Estimates are based on the 2004-2015 EU-SILC sample pooled accross countries. The sample consists of households formed by a one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had the first child at the earliest at the age of 16 and children's ages are in the range 0–12. The estimation method used is OLS. The explanatory variable is the dummy for a first-born daughter. The table presents estimated effects of the firstborn being a daughter compared the baseline case of the firstborn being a son. The dependent variables are the total number of children and a set of binary indicators for specific numbers of children. The control variables are: cubic polynomial in mother's age, squared polynomial in mother's age at first birth, length of cohabitation of spouses, mother's education, father's education, mother's employment, father's employment, household disposable income, living in urban area, tenure status, year and country dummies.

Table A3: Effects of the gender of the firstborn child in the household on the selected measures of fertility

Countries[a]	(1) Total number of children	(2) Two or more children	(3) Three or more children	(4) Four or more children	(5) Five or more children	Obs
AT	-0.0181	-0.0245*	0.0083	-0.0050	0.0015	6,574
BE	-0.0074	-0.0139	0.0054	0.0007	0.0004	7,694
BG	0.0206	-0.0112	0.0222**	0.0096*	0.0011	3,509
CH	0.0353	0.0364**	0.0013	0.0013	-0.0017	4,461
CY	-0.0422*	-0.0330**	-0.0125	0.0032	0.0002	5,675
CZ	-0.0123	-0.0167*	0.0037	-0.0002	0.0001	10,329
DE	-0.0141	-0.0179*	0.0060	-0.0012	-0.0010	9,790
DK	-0.0183	-0.0023	-0.0178*	0.0012	0.0007	7,889
EE	-0.0147	-0.0091	-0.0032	0.0027	-0.0017	6,594
EL	-0.0040	-0.0075	-0.0065	0.0045	0.0041***	8,147
ES	-0.0292**	-0.0277***	-0.0030	0.0003	0.0008	16,054
FI	-0.0027	-0.0031	0.0070	-0.0000	-0.0011	13,145
FR	0.0209*	0.0102	0.0072	0.0005	0.0029**	14,496
HR	0.0878**	0.0507*	0.026**	0.0127	0.0031	1,742
HU	-0.0082	0.0057	-0.0137**	-0.0027	0.0015	11,281
IE	0.0002	0.0094	0.0030	-0.0074	-0.0007	5,636
IS	-0.0059	0.0009	-0.0022	-0.0028	-0.0014	5,711
IT	0.0091	-0.0032	0.0121***	-0.0004	0.0002	21,486
LT	-0.0352	-0.0096	-0.0090	-0.0098**	-0.0040*	3,742
LU	-0.0068	-0.0069	0.0022	0.0020	-0.0029*	8,084
LV	-0.0172	-0.0204	-0.0020	0.0028	0.0008	5,102
MT	-0.0170	-0.0013	-0.0118	-0.0019	-0.0013	2,872
NL	0.0021	0.0039	-0.0033	-0.0001	0.0001	11,942
NO	-0.0385**	-0.0210*	-0.0191*	0.0006	0.0007	8,108
PL	0.0049	-0.0037	-0.0008	0.0023	0.0035**	18,374
PT	-0.0794	-0.0486***	-0.0216***	-0.0074**	-0.0008	6,044
RO	0.0293	0.0028	0.0218**	0.0075*	-0.0027	4,948
RS	0.0619	0.0378	0.0214	0.0044	-0.0017	1,221
SE	0.0240	0.0112	0.0114*	0.0019	-0.0006	9,228
SI	0.0140	-0.0147	0.0113	0.0093***	0.0036***	10,544
SK	0.0191	-0.0025	0.0093	0.0072*	0.0018	5,802
UK	-0.0155	-0.0104	0.0034	-0.0085*	-0.0012	9,288

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: See notes to Table 3 for data samples, variable definitions, and included control variables. The columns contain estimated country-level effects of the first-born in a household being a daughter on the corresponding variables in the column headings.

[a] Table A6 contains names of countries corresponding to the abbreviations.

Table A4: Abbreviations for countries

Abbrev.	Countries	Abbrev.	Countries	Abbrev.	Countries	Abbrev.	Countries
AT	Austria	EE	Estonia	IS	Iceland	PL	Poland
BE	Belgium	EL	Greece	IT	Italy	PT	Portugal
BG	Bulgaria	ES	Spain	LT	Lithuania	RO	Romania
CH	Switzerland	FI	Finland	LU	Luxembourg	RS	Republic of Serbia
CY	Cyprus	FR	France	LV	Latvia	SE	Sweden
CZ	Czech Republic	HR	Croatia	MT	Malta	SI	Slovenia
DE	Germany	HU	Hungary	NL	Netherlands	SK	Slovak Republic
DK	Denmark	IE	Ireland	NO	Norway	UK	The United Kingdom

Source: Eurostat

Table A5: Impact of the first-born daughter on selected household allocation decisions under two alternative explanations of the parental gender preference

Allocation decisions	Gender bias		Differential cost	
	Towards sons	Towards daughters	Higher outlays on sons	Higher outlays on daughters
Household public goods expenditure	-	+	.	.
Savings	+	-	-	+
Amount of money to make ends meet	+	-	+	-
Personal expenditures of a father	.	.	+	-
Personal leisure time of spouses	.	.	+	-
Availability of children's items at home	.	.	-	+

Note: The sign "+" means a positive impact and the sign"-" means a negative impact. The rationale behind the predictions is explained primarily in the Introduction and also in Sections 3 and 4.

Table A8: Spearman's rank correlations between country-level effects of the first-born daughter in the household on selected measures of fertility

	Total number of children	Progression to parity two	Progression to parity three	Progression to parity four	Progression to parity five
Total number of children	1				
Progression to parity two	0.8380***	1			
Progression to parity three	0.7878***	0.4765***	1		
Progression to parity four	0.4758***	0.2753	0.3680**	1	
Progression to parity five	0.0037	-0.1334	-0.0169	0.2834*	1

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: The Spearman's rank correlations are based on estimates for 32 European countries covered in EU-SILC survey during 2004-2015. The estimates are contained in the Table A5.

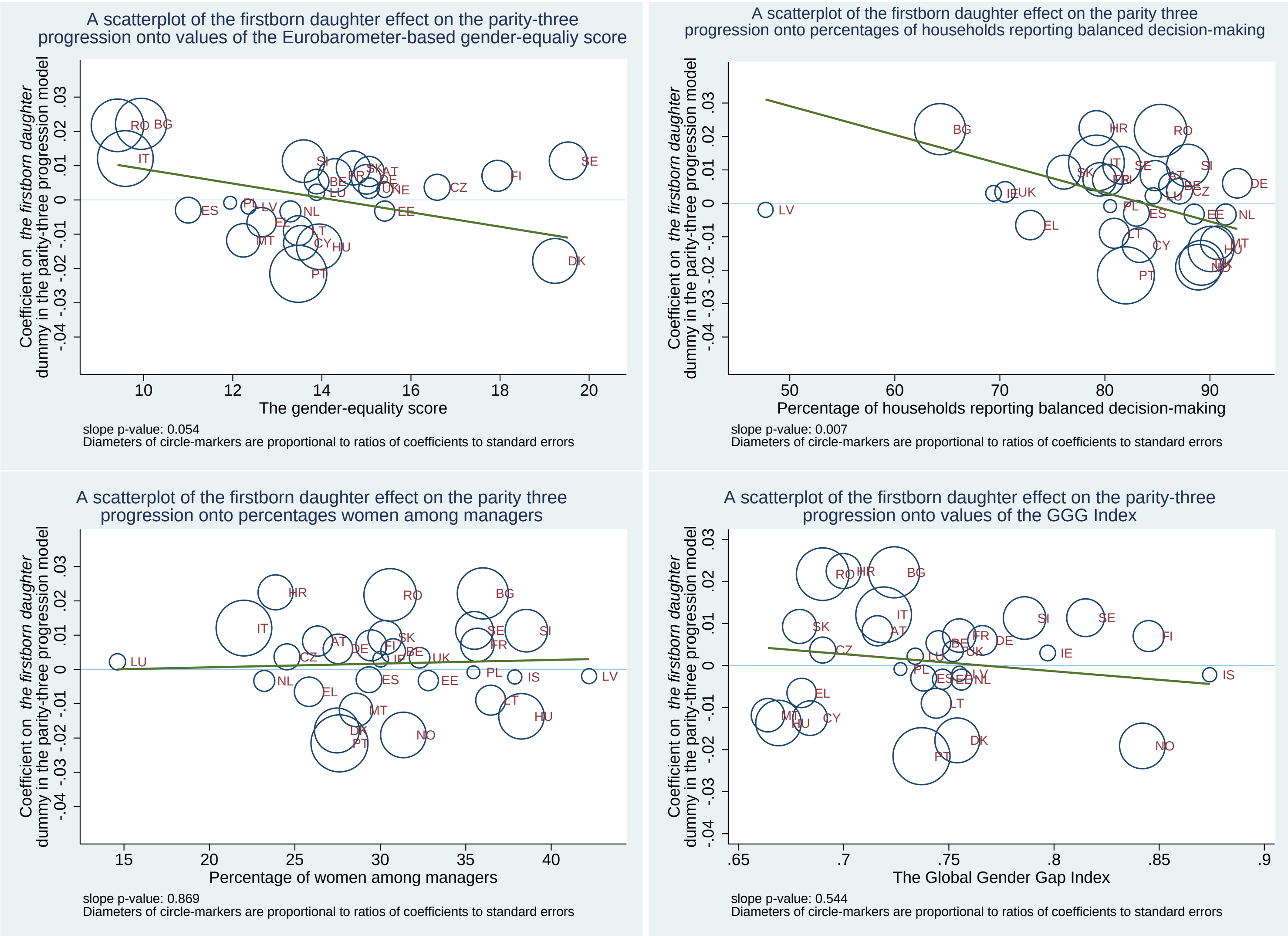


Figure A2: The relationship between the effect of the firstborn daughter on the third parity progression and specific gender-equality measures across countries. We calculate the Eurobarometer-based gender equality score for a particular country as a sum of this country's ranks in responses to questions about attitudes towards gender equality. These responses were collected in 2009 Eurobarometer special survey (Eur, 2010). For each question countries were ordered according to shares of respondents who perceive existence/wish to exist gender-egalitarian conditions in a specified realm of life. A country with the highest share of such respondents was assigned the rank 1 for the corresponding question. Then, we calculated sums of such ranks across all 13 pertinent questions and our gender-equality score. But, we do not have it for Switzerland, Croatia, Iceland, Norway, and the Republic of Serbia because the considered Eurobarometer survey was not conducted in these countries. Percentages of households reporting balanced decision-making were taken from the data of Health and Demographic Survey collected by the World Bank in different years and Survey of Income and Living Conditions collected by Eurostat in 2010. The percentage of women among managers was taken from the data of Enterprise Surveys conducted by the World Bank in different years. The Global Gender Gap Index was calculated by the World Economic Forum in 2016.

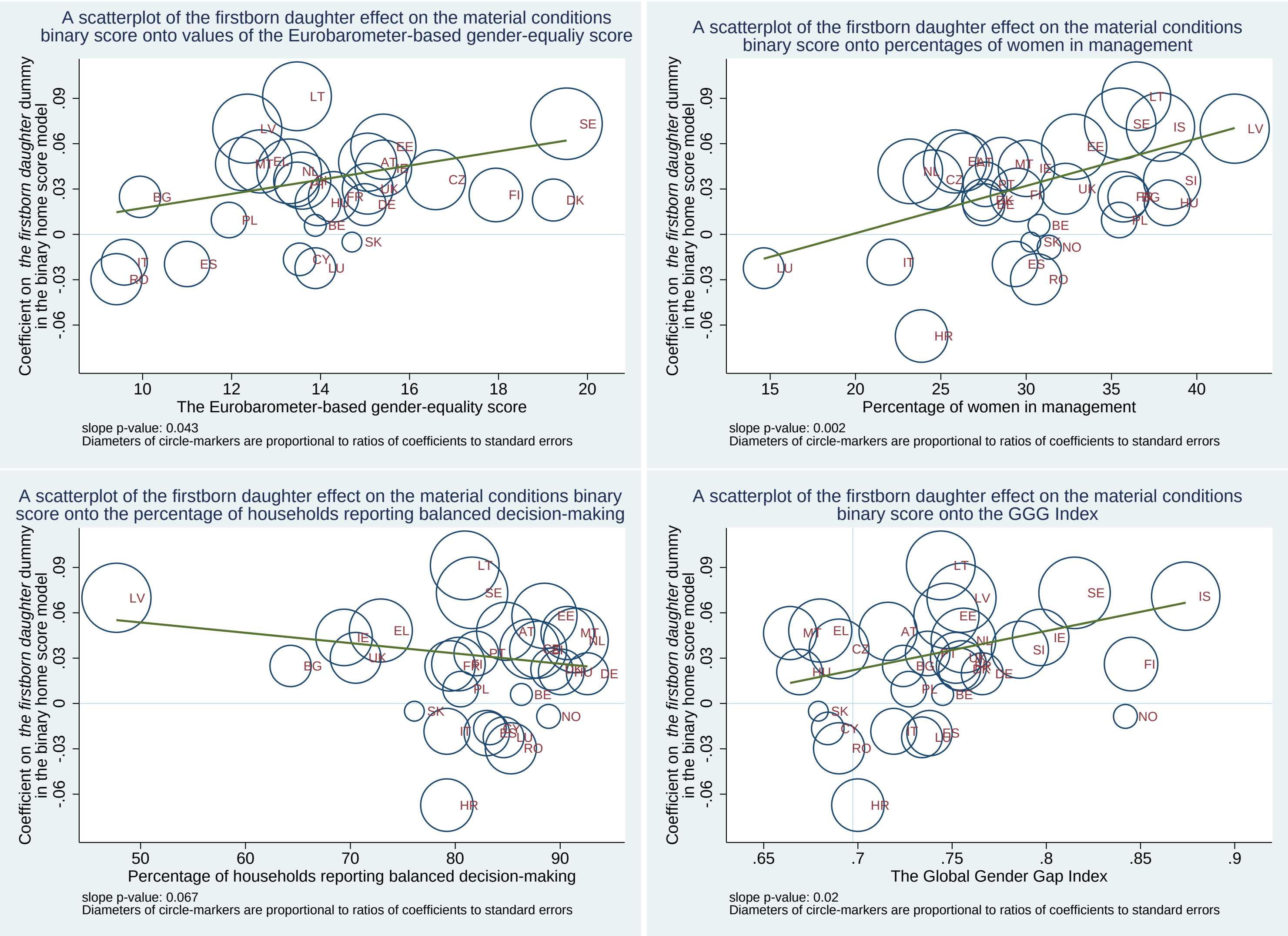
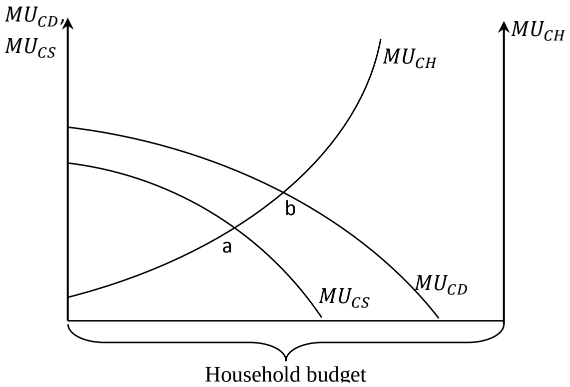
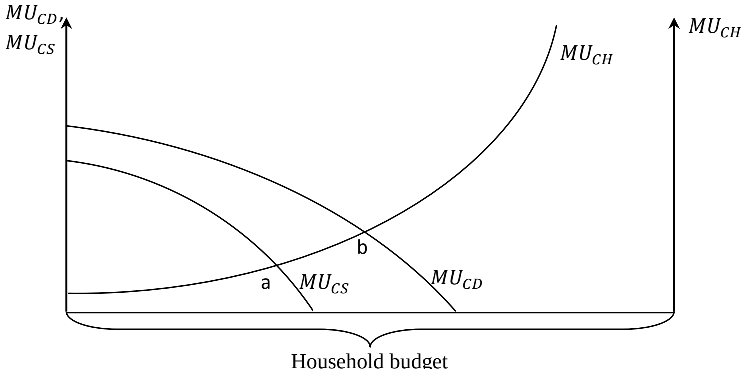


Figure A3: The relationship between the effect of the firstborn daughter on the children's material conditions and specific gender-equality measures across countries.



(a) Low-income families



(b) High-income families

Figure A4: The difference in expenditures on children between low-income and high-income households

Notes: See the note to Figure A1 for explantion.

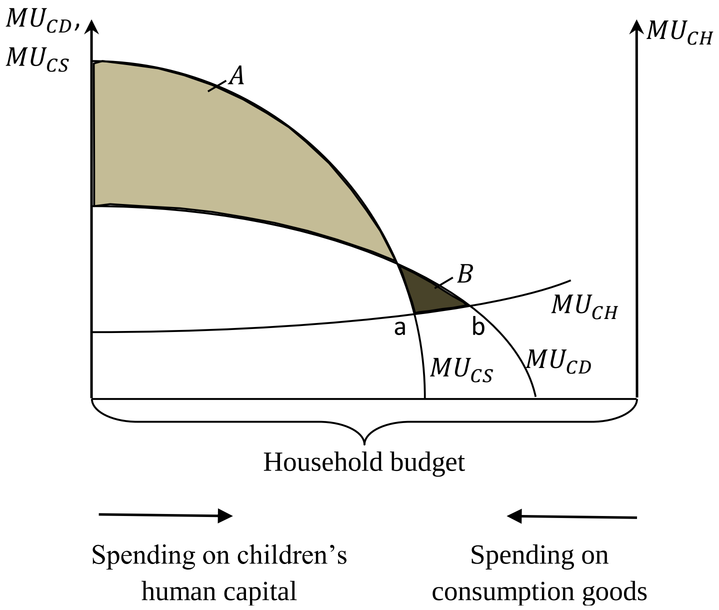


Figure A5: Coexistence of the son bias and the differential cost with prevailing effect of the son bias on fertility.

Notes: See the note to Figure A1 for explanation.

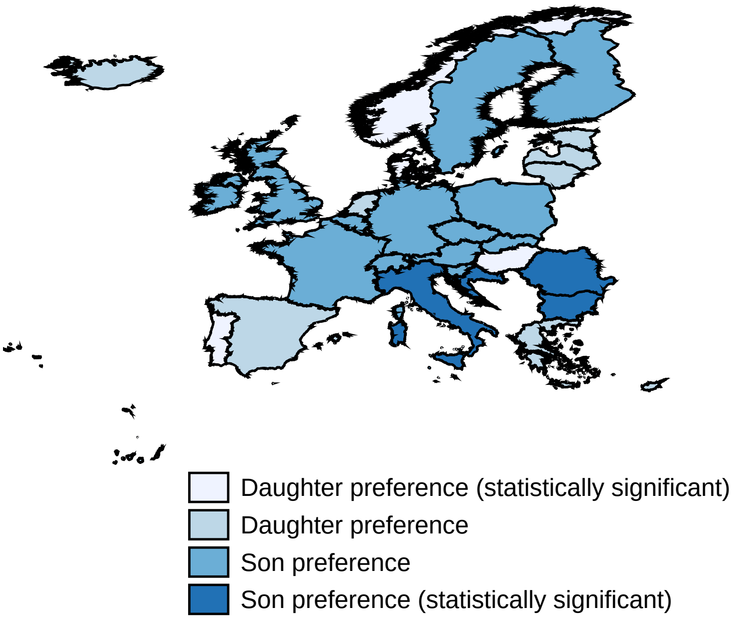


Figure A6: Gender Preferences for Children in 31 EU-SILC Countries

Table A9: The material deprivation indicator and the gender of firstborns.

	The binary material deprivation indicator on the LHS			
Explanatory var-s	(1)	(2)	(3)	(4)
	OLS	OLS	OLS	IV
First child a girl	.015***	.0148***	.0168***	.0172***
Number of children		.0896***	.0797***	-.0231*
Covariates	No	No	Yes	Yes
R-Square	.000225	.0191	.168	.146
N obs	51,087	51,087	49,922	49,922

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

Notes: The standard errors are clustered at the country level. Estimates are based on the 2004-2015 EU-SILC sample. The sample consists of households formed by a one cohabiting couple, their children, and, occasionally, other relatives. The mother of children in the household is younger than 41 and older than 17 and had the first child at the earliest at the age of 16 and children's ages are in the range 0–12. The estimation method used is OLS.The dependent variable is the binary indicators taking value 1 when a household has all 10 considered items listed in the Table 2 at disposal and value 0 otherwise. The control variables are: household disposable income, weekly work-hours of parents, dummies for parental employment status (full-time employment, self-employment, part-time employment), tenure status, living in urban area, ability to make ends meet, a woman responding the household questionnaire, parental educational attainment, and country. The estimates in the fourth column are obtained using 2SLS method from a regression-model in which the number of children is instrumented with the twin-birth. The first stage F-statistic value for this model is above two thousand.